The Impact of Exchange Rate Volatility on Turkish Exports: 1993-2009

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Abstract:

This paper attempts to investigate the long-run and short-run relationships between Turkish exports, exchange rate volatility, foreign income, and relative prices by employing quarterly data for the period 1993Q3-2009Q4. Towards this purpose, multivariate cointegration and error correction model (ECM) techniques are used in this study. The long-run estimation results suggest that foreign income and real exchange rate volatility exert positive and statistically significant impacts on Turkish exports, while relative prices affect Turkish exports negatively and significantly. In addition, the results of the ECM model indicate that relative prices have a negative and significant effect, foreign income has an insignificant effect, and nominal exchange rate volatility has a positive and significant effect on Turkish exports.

Keywords: Exchange Rate Volatility, Export, ARDL Cointegration, Turkey.

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1. Introduction

Since the collapse of the Bretton-Woods system, the volatility of real and nominal exchange rates has increased among countries that adopted a new regime of floating exchange rates. In addition, free movement of capital cross-border also tended to amplify the extent of exchange rate volatility among countries. As a result of these developments, many researchers have shifted their attentions to investigating the nature and magnitude of the impact of exchange rate volatility on trade flows. Therefore, a great deal of both theoretical and empirical literature has been generated in this field. From this literature two distinguished hypotheses have emerged: whilst the first argues that the volatility of exchange rates would have an adverse effect on trade flows, the second argues that the volatility of exchange rates would encourage trade flows. In the Turkish context, information about the impact of volatility on exports is also a major concern for policy makers given the implementation of flexible exchange rate mechanism since early 2001.

Empirical studies have also yielded conflicting evidence for the effect of exchange rate volatility on trade flows due to the sample period chosen, model specification, measure of volatility preferred, and countries selected. There are a great number of empirical studies that support the hypothesis that increased

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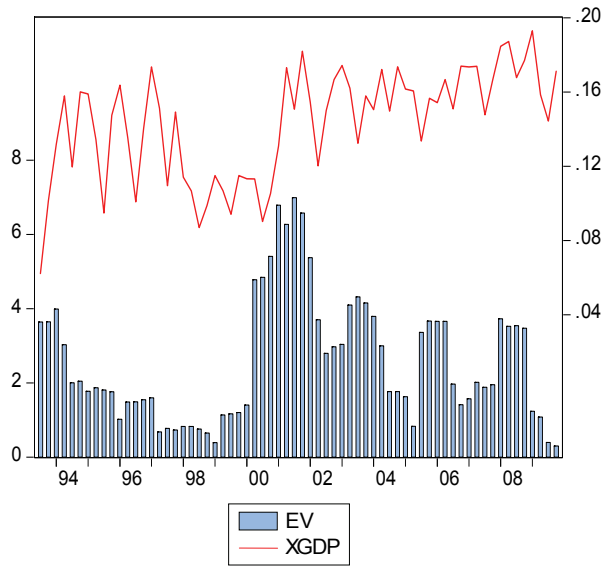
exchange rate volatility will decrease trade flows due to risk-averse traders. Among the examples are Cushman (1983 and 1988), Kenen and Rodrick (1986), Thursby and Thursby (1987), De Grauwe (1988), Chowdhury (1993), Arize (1995), Doganlar (2002), Vita and Abbot (2004), Clark, et al, (2004); and Arize et al (2008). On the other hand, several others support the hypothesis that increased volatility may lead to a greater volume of trade or an ambiguous effect on trade flows (Hooper and Kohlhagen, 1978; Bailey et al, 1987; Franke, 1991; Asseery and Peel, 1991; Sercu and Vanhulle, 1992; Holly, 1995; McKenzie and Brooks, 1997; McKenzie, 1998; Aristotelous, 2001; Bredin et al, 2003; Kasman and Kasman, 2005; and Zhang et al, 2006).

The effect of volatility in exchange rates is particularly important for Turkish exports since more than 25 percent of GDP during the study period is exported to the world market. The results of the current estimation are to some extent consistent with the previous findings by Kasman and Kasman (2005) and Ozturk and Kalyoncu (2009). The earliest work of Ozbay (1999) covering this period, where a crawling peg exchange regime was followed, concluded that exchange rate uncertainty bears a negative impact on exports. Another study by Vergil (2002), estimating bilateral exports to the US, Germany, France, and Italy also provided evidences for a negative relationship between real exchange rate volatility and real exports. Kose et al (2008), using three different measures of real exchange rate volatility over the monthly period 1995-2008, found that exchange rate volatility has a negative effect on aggregate exports both in the long-run and short-run. Other studies which also established similar findings include Buguk et al (2001), Doğanlar (2002), Ozturk and Acaravcı (2003), Demirel and Erdem (2004), Saatcioğlu ve Karaca (2004), and Tuncsiper and Oksuzler (2006).

As can be seen from Figure 1, higher exchange rate volatilities in 1994, 2001, and 2008 are associated with greater instabilities in the economy and world as well as significant increases in scale and variety of cross-border financial transactions in the last two decades. In addition, real exchange rate has become fairly volatile since Turkey followed free float exchange policy after 2001 compared with the period before 2001 when the managed float and crawling peg exchange regimes were adopted. If the exchange rate fluctuations, particularly for the post-crisis period (2001), are examined together with the export performance of Turkey, one can suggest that the export performance does not fall with increasing exchange rate fluctuations.



Note: Real Exchange volatility is computed by using data from the Central Bank of Turkey

Figure 1: Real Exchange Volatility and Export/GDP Ratio, 1993-2009

The rest of the paper is organized as follows: Sections 2 and 3 review the previous theoretical and empirical literature, respectively; Section 4 specifies the model of export demand of Turkey, describes data sources and definitions of variable and explains the estimation techniques. Section 5 discusses the findings of empirical analysis. Finally, section 6 offers some concluding remarks and policy implications.

2. Theoretical Literature

As was said earlier, the impact of exchange rate volatility on trade flows can be positive or negative depending on the assumptions employed in relation to the definition of exchange rate volatility, market structure, the flexibility of production capacity, the presence or absence of a forward exchange market, and the risk preferences of traders (Hooper and Kohlhagen, 1978; Ethier, 1973; De Grauwe, 1988; Franke, 1991; Viaene and Vries, 1992; Sercu and Vanhulle, 1992; and De Grauwe, 1994). Côte (1994) and McKenzie (1999) who surveyed the literature on the topic conclude that there has been no unambiguous relationship between exchange rate volatility and trade flows.

The early theoretical studies supported the view that an increase in exchange rate volatility leads to a decline in

the volume of trade flows (Clark, 1973; Hooper and Kohlhagen, 1978). These models, by taking into account risk-averse firms, argued that higher exchange rate volatility lowers the level of expected profits from exports and, thus reduces the volume of trade. This outcome is obtained with the assumptions of no forward market, that exchange rate volatility is the only source of risk, and production and trade decisions are to be made before uncertainty is resolved. Another traditional model of Ethier (1973) and Baron (1976), under the assumption of perfect hedging, stated that exchange rate volatility may not affect the volume of trade. This is the case for industrial countries, where there is a well-developed forward market. Firms in these countries can easily avoid their sales from markets where there is high exchange rate volatility. In addition, there are various possibilities for reducing the level of exposure to the exchange rate risks for a multinational firm that engages in a wide variety of trade and financial transactions across a large number of countries (Makin, 1978).

On the other hand, recent studies suggest no clear-cut relationship between exchange rate volatility and trade flows. Viaene and Vries (1992) stressed that even in the presence of a forward market, exchange rate volatility can affect indirectly the volume of trade. They show that an increase in exchange rate volatility may generate adverse effects on importers and exporters since they are located on the opposite sides of the forward contract. Thus, an increase in exchange rate volatility may operate to the loss or benefit for trade flows, depending on the net trade balance of the country. Exporters lose (benefit) and importers benefit (lose) when the trade balance is positive (negative), or alternatively when the forward risk premium is positive (negative). De Grauwe (1988), assuming firms operate in a competitive market and domestic price of exports is the only source of risk, suggested that the effect of exchange rate volatility on exports depends on the degree of risk aversion. If firms are sufficiently risk averse, an increase in the exchange rate volatility raises the expected marginal utility of export revenue and thus induces them to export more (income-effect). However, if firms are less risk averse, they will produce less for exports since exporting becomes relatively less attractive (substitution-effect). Under this setting, the author concluded that the dominance of the income-effect over the substitution-effect can lead to a positive relationship between exchange rate volatility and trade flows.

While these models took into account the firm's option to hedge or at least diversify its exchange risk, they often ignored the firm's option to adjust producton and trade in response to exchange rate risks. Models focusing on such adjustments suggested that higher exchange rate volatility can cause more trade (see Dixit, 1989; Franke, 1991; Sercu and Van Hulle, 1992; De Grauwe, 1994). Franke (1991) modelled the export strategy of a risk neutral firm, which operates in a monopolistic market. The export strategy is determined by the transaction (entry and exit) costs to a foreign market. A firm starts exporting if the present value of the transaction costs is outweighed by the present value of the expected cash flows from exports. When the expected cash flow is a convex function of the exchange rate, the present value of cash flows grows faster than that of the entry and exit costs and the firm benefits from increased exchange rate risk. Under this setting, the model predicts that firms will enter a market sooner and exit later when the exchange rate volatility increases, so that the amount of firm trade will increase.

Sercu and Vanhulle (1992) analyzed the behavior of an exporter who incurs standard duties and transport costs instead of the entry and exit costs to/from a foreign market. In their analysis, Sercu and Vanhulle assumed risk aversion, a perfect forward market and a random walk for the exchange rates. When exchange rate falls (appreciation of local currency) below a certain level, the firm can either abandon the market completely or suspend exports temporarily. In the first case, all expenditures are stopped, while in the second case, the firm continues to incur some costs – such as maintaining equipment and keeping in contact with the market. The results of their analysis indicated that an increase in exchange volatility raises the value of the exporting firm and reduces the exchange rate, which causes the firm to stop exporting completely.

Finally, De Grauwe (1994) developed a very simple static model of an exporting firm which is a price taker in its market. The model assumed that there are no adjustment costs and the firm responds rationally to changes in production prices. A weaker (stronger) exchange rate increases (decreases) the local currency prices of exports, which induces the firm to expand (contract) production. Given the lack of consensus on the theoretical effects of exchange rate volatility on exports, there is a greater need for empirical analysis to shed light on this matter.

3. Empirical Literature

Similar to the theoretical discussions, empirical evidences of both early and recent studies have mixed implications, with a large majority of these studies failing to produce systematic relationship between volatility and trade flows. The early studies surveyed by the IMF (1984) did not yield consistent results, with a large majority producing little or no support for a negative relationship between exchange risks and trade flows, while bilateral studies seem to establish a negative relationship between the two variables. The earliest work conducted by Hooper and Kohlhagen (1978) examined the impact of exchange rate volatility on aggregate and bilateral trade flows for all G-7 countries except Italy. They measured the exchange rate volatility by the absolute difference between the current period spot exchange rate and the forward rate of the last period, as well as the variance of nominal spot exchange rate and the current forward rate. They found basically no evidence of any negative relationship.

Cushman (1983) used real exchange volatility to estimate U.S. bilateral exports to 17 industrial countries over the quarterly period of 1965-1977. From the 28 equations estimated in total, he found seven negative and five positive and significant coefficients for exchange rate volatility. Another study carried out by the same author in 1988 explored the relationship between real exchange rate volatility and U.S. bilateral trade flows. The study covered the period 1974-1983, where floating exchange regime was adopted. The results revealed a large number of significant and negative risk effects in 5 of 6 import flows and only one positive and significant effect on exports.

The IMF (1984) estimated bilateral export functions between G-7 countries over the quarterly period of 1969-1982. Export is considered as functions of foreign income, capacity utilization, real exchange rate, and real exchange volatility, which was measured as the standard deviation of percentage changes in the exchange rates over the preceding five quarters. From a total of 42 estimations made, only three provided negative and significant relationships, while eleven cases provided positive and significant relationships between the exchange volatility and exports.

Asseery and Peel (1991) also examined the influence of effective exchange rate volatility on the multilateral exports of five industrial countries over the quarterly period of 1972-1987. They used cointegration and error correction techniques to explore the long-run and dynamic short-run effects of exchange volatility on exports. For all countries except for the UK, the study found that volatility has positive and significant effects on exports.

A study by Kroner and Lastrapes (1993) examined the impact of exchange rate volatility on the exports of five developed countries for the monthly period of 1973-1990, using multivariate GARCH models. The results of the study indicated that exchange rate volatility have statistically significant effects on the export volumes for all countries. The signs and magnitudes of the volatility coefficients differ widely across the countries; negative for the U.S. and U.K. and positive for Germany, Japan, and France.

McKenzie and Brooks (1997) used the ARCH model to quantify variability in exchange rates in their study of U.S.-German bilateral trade over the monthly period 1973-1992. In their model, trade is viewed as a function of income, relative prices, exchange rates, and volatility. The estimation results indicated that a significantly positive relationship exists between volatility and trade flows.

Arize et al. (2000 and 2008) estimated export models to examine the effect of real exchange rate variability on export flows of 13 less developed countries and 8 Latin American countries, respectively. They used Johansen's multivariate cointegration and error correction techniques over the quarterly periods 1973-1996 and 1973-2004 to measure the long-run and short-run dynamic impacts of exchange rate volatility on exports. The major results showed that increases in exchange rate volatility impose significantly negative effects on exports both in the long-run and in the short-run in all countries.

Bredin et al. (2003) analyzed the relationship between export volume and its determinants, by using cointegration and error correction techniques. The model was estimated for both aggregate and sectoral Irish exports to the EU over the quarterly data for the period 1978-1998. The short-run and long-run estimation results showed that exchange volatility has no significant effect in the short-run, but has a positive and significant effect in the long-run on aggregate and sectoral exports. This can be concluded that a decline in exchange volatility associated with single currency (euro) will lead to a long-run fall in Irish exports to the market.

Aurangzeb et al. (2005) examined the impact of exchange rate volatility on Pakistan's exports to its major trading partners over the monthly period 1985-2001. Long-run and short-run export models were estimated by utilizing Johansen's multivariate cointegration and errror

correction approaches. Exchange rate volatility is measured by using an ARCH-type model. The long-run results provided a negative and insignificant impact of volatility on exports in all cases. The results of the short-run models, however, suggested a negative causality running from volatility to exports in all cases except for the UK.

Hall et al (2010) examined the impacts of exchange rate volatility on exports of 10 emerging market economies (EMEs) and 11 other developing countries in their study. They used panel data, covering the period 1980:Q1–2006:Q4 and two different estimation methods - generalized method of moments (GMM) and a time varying coefficient (TVC). The result suggested that there is a negative and significant effect of exchange volatility on exports for the non-EMEs, while there is no negative effect for the EMEs. This implies that the open capital markets of EMEs may have reduced the effects of exchange rate fluctuations on exports compared with those effects in the other developing countries.

There are also studies that employ an augmented version of the gravity model[1] to investigate the relationship between exchange rate volatility and trade flows. These studies also failed to establish any systematic relationship between the two variables. Frankel and Wei (1993) looked into the relationship between exchange rate volatility and bilateral trade flows between 63 industrial and developing countries by using cross section data for 1980, 1985, and 1990. The empirical results found that exchange rate volatility has negative and significant effects on trade in 1980 and positive and significant effects in 1990.

Dell' Ariccia (1999) investigated the impact of nominal and real exchange rate volatility on bilateral trade between 15 EU countries and Switzerland over the period 1975-1994. He used four different measures of volatility and found negative, but small (ranging between -0.10 to -0.13) coefficients on the exchange rate volatility, implying that an elimination of the exchange rate volatility would raise bilateral trade between Switzerland and the 15 EU countries by 10 to 13 percent. When fixed and random effects are incorparated in the estimation procedure to explain the simultaniety bias, however, the effect of exchange volatility reduces significantly to 3-4 percent.

[1] The gravity model has been used widely in empirical works and has been highly successful in explaining bilateral trade flows. In its basic form, the gravity model explains bilateral trade flows as functions of their GDPs, geographical distance, population, and several dummy variables, such as common border, common languages, and membership in a free trade organization (Clark, 2004: 6).

Another study, Rose (2000), which also employed the gravity approach, used a data set covering 186 countries for the five years 1970, 1975, 1980, 1985, and 1990 to test the impact of exchange rate volatility on trade. The estimation results also confirmed a negative relationship between exchange rate volatility and bilateral trade, with the volatility effect being relatively small (a coefficient of -0.13). This implied that reducing exchange rate volatility by one standard deviation around the mean would increase bilateral trade by 13%. However, when random effects are taken into account in the estimation procedure, the size of the effect of exchange volatility falls significantly to 4 percent.

Tenreyro (2003) used a panel data set covering a period from 1970 to 1997 for a large number of countries to address the relationship between exchange rate volatility and bilateral trade. This study carried out different estimation methods to overcome statistical problems that were faced by the previous studies and obtained values for the coefficients of exchange volatility ranging between 2 and 4 percent. This implied that eliminating all exchange rate volatility would increase trade by a maximum of 4 percent. However, when endogeneity is taken into consideration by the use of instruments, volatility seems to lose its significant effect on trade.

4. Model Specification and Data Set

4.1. Model Specification

Based on the early empirical studies of exchange rate volatility and following Arize et al. (2000), Bredin et al. (2003), and Aurangzeb et al. (2005), the long-run export demand equation can be written in log-linear form as:

$$X_t = \beta_0 + \beta_1 Y_t^* + \beta_2 PI_t + \beta_3 EV_t + \beta_4 D_1 + \beta_5 D_2 + u_t \quad (1)$$

where X_t represents real exports (measured as the ratio of nominal exports to nominal GDP); Y_t^* is a measure of foreign economic activity, which is proxied by GDP deflator of OECD countries; PI_t represents relative prices (defined as the ratio of Turkish export prices to world export prices); EV_t is a measure of exchange rate volatility, which is measured by the moving average of the standard deviation of real effective exchange rates;

and D_1 and D_2 represent dummy variables that took into account the possible effects of the Russian crisis of 1998 and the Turkish financial crisis of 2000. Economic theory suggests that increases in the real GDP of trading partners would result in a greater volume of exports to those partners. However, if the relative prices rise (fall), it would make domestic goods less (more) competitive than foreign goods and, therefore, the demand for exports will fall (rise). As discussed in the previous section, the effect of exchange rate volatility on exports is ambiguous. Therefore, it is expected that $\beta_1 > 0$, $\beta_2 < 0$, and β_3 is either positive or negative.

Exchange rate volatility is a measure that aims to capture the risks faced by exporters due to unexpected fluctuations in exchange rates. Various volatility measures have been proposed in the literature, including averages of absolute changes, standard deviations, deviations from trend, and variances of exchange rate changes. In this study, we implimented the most commonly employed method, namely the moving average of the standard deviation of real effective exchange rate changes, since it is more likely to account for periods of high and low exchange rate uncertainity. This proxy is mathematically expressed as:

$$EV_t = \left[\frac{1}{m} \sum_{i=1}^{m} (\ln ER_{t+i-1} - \ln ER_{t+i-2})^2 \right]^{\frac{1}{2}}$$

where R represents the real effective exchange rates and m is the order of moving average and equals 4 in this work.

4.2. Data and Methodology

The study covers quarterly observations from 1993:3 to 2009:4. The data on exports, Turkish exports price index, and Turkish GDP are obtained from International Financial Statistics available on the internet sites of the IMF. The GDP deflator of OECD countries is collected from the online sources of the OECD website, while real effective exchange rate is obtained from the Turkish Central Bank. Finally, world export price index is constructed with the use of the export price index of nine major trading partners, which is available from the sources of Eurostatistics. All real values are measured through the base of year 2000. All these series are expressed in US dollar terms except for the Turkish GDP, which is then converted into US dollars by using bilateral US/TL exchange rates. All of the series, except for EV, were transformed into logarithmic forms in order to interpret the results as elasticities. The export series was the only variable that required a seasonal adjustment process and it was accomplished with the Troma/Seat method with the econometric software Eviews.

Prior to estimating eq. (1), the time-series properties of the individual variables must be checked to ensure that none of the variables are stationary at their second differences, I(2), since most of the time series tend to be non-stationary and their variances increase over time in level. The estimation of a model with such variables (integrated of order 2) tends to produce spurious results. Here the order of the integration of the above variables is determined using the augmented Dickey-Fuller (ADF) and Phillips-Peron (PP) unit root tests by the following formulas;

Variables		Level		First Difference		Results
		ADF	PP	ADF	PP	
X		-3.17 (0)**	-3.21 (1)***	-6.68 (3)***	-8.32 (8)***	I (0)
Y*		0.41 (1)	0.88 (4)	-4.27 (0)***	-4.27 (0)***	I (1)
PI		-3.34 (0)**	-3.32 (3)**	-4.14 (0)***	-10.8 (10)***	I (0)
EV		-2.62 (10)	-2.29 (4)	-4.95 (3)***	-7.02 (3)***	I (1)
Critical Values	%1	-3.53	-3.51	-3.53	-3.51	
	%5	-2.90	-2.89	-2.90	-2.89	

Note: Figures in parentheses represent the number of lags chosen with respect to the AIC. *** and ** denote the rejection of null hypothesis (that variables are non-stationary) at the 1% and 5% significance level, respectively. The critical values for ADF and PP tests are obtained from MacKinnon (1996). Estimations are carried out by Eviews econometric software.

Table 1: Unit Root Test Results

$$\Delta y_t = \alpha + \delta y_{t-1} + \sum_{i=1}^{k} \gamma_i \Delta y_{t-i} + \varepsilon_t$$

where Δy is the first difference of y series, α is a constant term, ε_t is the residual term and k is the lagged values of Δy_t,which is included to avoid serial correlation in the residuals. The ADF and PP test results without trend are provided in Table 1. The lag length for the ADF tests was selected based on Akaike Information Criteria (AIC), with the estimation of an initial eleven lagged first-differenced right-hand side variable, Δy_{t-i}. The results of the integration tests clearly show that foreign income and EV series are non-stationary in levels, but stationary in first differences, I(1) at the 1% significance level, while the other two series are level stationary. Hence, the results of these integration tests support the use of the Autoregressive Distributed Lag (ARDL) approach in this paper.

In terms of methodology, this paper adopts the ARDL cointegration technique developed by Pesaran and Shin (1999) and Pesaran et al (2001) to establish the long-run relationship between the variables. This estimation procedure is preferred to other single multivariate cointegration tecniques for four reasons. First, the ARDL approach does not require unit root tests for the variables in the model and can be applied to any set of data, which are purely I(0), purely I(1) or a mixture of both. Second, endogeneity problems and the inability to test hypotheses on the estimated coefficients in the long-run associated with the Engle-Granger (1987) method are avoided. Third, with the ARDL approach, it is possible for different variables to have different optimal lags, which is impossible with standard cointegration techniques. Finally, this procedure is more efficient than the others with limited sample data (30 to 80 observations), as suggested by Narayan (2005).

Basically, the procedure involves estimating the error correction version of the ARDL model for eq. (1) and can be formulated as follows:

$$\Delta X_t = a_0 + \sum_{i=1}^{m} a_{1i} \Delta X_{t-i} + \sum_{i=0}^{m} a_{2i} \Delta Y^*_{t-i} + \sum_{i=0}^{m} a_{3i} \Delta PI_{t-i} + \sum_{i=0}^{m} a_{4i} \Delta EV_{t-i} + a_5 X_{t-1} + a_6 Y^*_{t-1} + a_7 PI_{t-1} + a_8 EV_{t-1} + \lambda_1 D_1 + \lambda_2 D_2 + u_t \quad \textbf{(2)}$$

where m stands for the lag length; u_t is white noise error terms; and the first line starting with the summation letter, Σ, represents the short-run dynamic relationship, while the second line (one-year lagged variables) represents the long-run relationship.

In order to determine whether a long-run relationship exists among the variables in hand we must first test the null hypothesis of no cointegration, ($H_0 : a_5 = a_6 = a_7 = a_8 = 0$) against the alternative hypothesis of cointegration, ($H_1 = a_5 = a_6 = a_7 = a_8 = 0$) using the F-test for the joint significance of the coefficients of the lagged levels in equation (2). The critical value of the F-test, which has a non-standard distribution, depends on (i) the integration level of the data, (ii) the number of regressors, and (iii) whether the model contains an intercept and/or a trend. Therefore, two sets of critical values are generated, one set referring to the I(0) series (lower bound critical values) and the other to I(1) series (upper bound critical values). If the computed F-statistic is greater than the upper bound critical value, then the H_0 hypothesis is rejected. If the computed F-statistic is less than the lower bound critical value, it implies no cointegration. Finally, if it lies between the lower and upper bounds critical values, then the result would be inconclusive. In the second step, once a cointegration is established among the variables, the following long-run export function can be written as follows:

$$X_t = b_0 + \sum_{i=1}^{m} b_{1i} X_{t-i} + \sum_{i=0}^{m} b_{2i} Y^*_{t-i} + \sum_{i=0}^{m} b_{3i} PI_{t-i} + \sum_{i=0}^{m} b_{4i} EV_{t-i} + b_5 D_1 + b_6 D_2 + u_t \quad (3)$$

The order of lag (m) in eq. (3) is selected using Akaike Information Criteria (AIC) before the selected model is estimated by ordinary least squares. Since we use quarterly data, our estimations start with a maximum of 6 lags and choose the model that minimizes the value of AIC (1, 0, 0, 1). In the third and final step, we obtain the short-run dynamic paramaters by estimating an error correction version of the ARDL (4, 0, 2, 0) model. This is specified as follows:

$$\Delta X_t = c_0 + \sum_{i=1}^{m} c_{1i} \Delta X_{t-i} + \sum_{i=0}^{m} c_{2i} \Delta Y^*_{t-i} + \sum_{i=0}^{m} c_{3i} \Delta PI_{t-i} + \sum_{i=0}^{m} c_{4i} \Delta EV_{t-i} + c_5 EC_{t-1} + e_t \quad (4)$$

where c_5 represents the error correction coefficient which measures the speed of adjustment to the long-run equilibrium. A negative and significant coefficent of EC_{t-1} will be an indication of cointegration.

5. Empirical Results

Before estimating the long-run and short-run export models for the Turkish economy we must test the null

hypothesis of no cointegration among the variables. In doing this, we first decide the lag order on the first differenced variables in eq. (2) by using AIC. The optimal lag length (m=1) is chosen to reveal that there is no serial correlation among the residuals obtained from eq. (2). The results of the ARDL bounds tests with the number of lags in row 1 are provided in Table 2.

Lag Order	LM tests	AIC	SBC	F-tests
6	1.44 (0.26)	-1.78	-0.58	1.77
5	5.31 (0.01)***	-1.80	-0.76	1.99
4	5.10 (0.01)***	-1.75	-0.85	1.73
3	0.29 (0.74)	-1.68	-0.93	5.23***
2	1.11 (0.34)	-1.67	-1.06	5.71***
1	1.40 (0.49)	-1.85	-1.38	8.36***

Note: *** denotes significance at the 1% level. LM represents the lagrange multiplier test for serial correlation. Number inside the brackets represents the probability ratio. F-statistics for 1% critical value bounds are 3.74 and 5.06. The critical values are obtained from Table CI (iii) in Pesaran et al (2001). Estimations are carried out by Eviews econometric software.

Table 2: Bounds Tests for Cointegration

The computed F-statistics (8.36) is greater than the upper bound critical value (5.06) at the 1% significance level, using an unrestricted intercept and no trend. Thus, the null hypothesis of no cointegration is rejected, implying long-run cointegrated relationship between Turkish exports and its determinants. The empirical results of the long-run cointegration model obtained by normalizing on Turkish exports are provided in Table 3. As can be observed from the table, all the variables, ($Y^*, PI,$ and EV) appear to have significant effects on the dependent variable, (X) and their coefficients have the expected signs. Thus, this result seems to be consistent with the existing empirical literature. Our results can be interpreted in the following way;

Dependent Variable	Independent Variables		
X	Y^*	PI	EV
	0.84 (3.36)***	0.34 (0.85)	0.04 (2.38)**

Note: *** and ** denote significance at the %1 and %5 levels. Numbers in parantheses represent t-values. Estimations are conducted with the use of Eviews econometric software.

Table 3: Estimated Long-run Coefficients with the ARDL (1, 0, 0, 1) Model

First of all, a positive and significant coefficient on foreign income variable is consistent with our expection and previous studies. This means that a rise in the level of OECD income leads to an increase in the level of Turkish exports. A 10% rise in the level of foreign income will cause an 8.4% increase on Turkish exports. The relatively low income elasticity, compared to the previous studies, is probably linked to the use of OECD GDP rather than world GDP. Many studies carried out for Turkey, at aggregate level, have found a long-run income elasticity that was greater than one (Acaravcı and Ozturk, 2003; Kasman and Kasman, 2005; Tuncsiper and Oksuzler, 2006; Kose et al, 2008; Ozturk and Kalyoncu, 2009).

Second, a positive but statistically insignificant coefficient on the relative export price variable implies that a change in the relative prices is not related to a change in the level of Turkish exports. Therefore, given the price inelasticity of exports, Turkey cannot use price competition policies to maintain or even increase its world's export share rather than adopting alternative policies.

Third, and most importantly, exchange rate volatility has a positive sign and affects the value of Turkish exports significantly. This finding is consistent with the previous studies of Kasman and Kasman (2005) and Ozturk and Kalyoncu (2009) in the case of Turkey. There are also overseas studies that support the positive relationship between volatility and exports, such as Klein (1990) and McKenzie and Brook (1997). Exporters increase the supply of exports in response to changes in exchange rate volatility when such volatility (which is 2.58% in the sample period) is above the threshold level (see, Zhang et al, 2006). As was argued before by Franke (1991), trade would increase with exchange volatility when the expected cash flow from exports grows faster than the entry and exit costs.

The final stage in this work is to construct an error correction model (ECM). The model structure is determined by Hendry's general to specific model

Variable	Coefficent	t –statistic	Variable	Coefficient	t -statistic
C	-0.005	-0.28	ΔPI_t	0.29	1.07
ΔX_{t-1}	0.36	1.86*	ΔPI_{t-1}	-0.30	-1.22
ΔX_{t-2}	-0.09	-0.85	ΔPI_{t-2}	0.37	1.52
ΔX_{t-3}	-0.09	-0.84	ΔEV_t	-0.003	-0.26
ΔX_{t-4}	-0.27	-2.42**	EC_{t-1}	**-0.74**	**-3.11*****
ΔY_t^*	3.88	0.80			
Diagnostic Test Results					
F-statistic = 3.08 (0.004)			$\chi^2_{BG}(2)$ = 0.38 (0.68)		
DW-statistic = 2.07			ARCH F(1, 58) = 1.12 (0.29)		
Adjusted R^2 = 0.38			RESET F(1, 49) = 2.06 (0.16)		
AIC = -1.80			χ^2_{JBN} = 28.3 (0.000)		

Note: The optimal lag order is determined by AIC. Numbers in parentheses show p-values. ***, ** and * denote significance at the %1, %5 and %10 levels. χ^2_{BG}, ARCH, RESET, and χ^2_{JBN} show Breusch-Godfrey test statistics for autocorrelation, heteroskedasticity, Ramsey misspecification, and Jarque-Bera normality statistics.

Table 4: Estimated Short-run Coefficients with the ARDL (4, 0, 2, 0)

selection strategy. This procedure requires elimination of insignificant lags from the estimation of eq. (4). The regression results are reported in Table 4. Before discussing the results, we need to determine the consistency of the ECM. For this reason, we performed a number of diagnostic tests that are reported at the bottom of Table 4. These tests indicate that the short-run model has no serial correlation, heteroskedasticity, and mis-specification problems. The following remarks can be made from the estimation results.

First, in contrast with the long-run results, the short-run income coefficient is positive, but statistically insignificant. Second, unlike the long-run estimation results, the short-run exchange rate volatility coefficient is negative and statistically significant. Third, the error correction term, EC_{t-1}, is negative and statisitically significant at the 1% level, indicating that the adjustment of short-run fluctuations towards the long-run equilibrium point takes place very fast.

5. Conclusion and Policy Implications

The purpose of this paper was to examine the impact of real exchange rate volatility on aggregrate Turkish exports over the quarterly period of 1993:3 to 2009:4 by employing the ARDL cointegration approach and error correction model. There are three important findings in this study. First, the results of the unit root tests indicate that variables under consideration are either I(0) or I(1) in level, suggesting that the ARDL approach is more appropriate. Second, the estimation results from the long-run model suggest that Turkish exports are significantly related to foreign income and exchange rate volatility, while they are not significanty associated with relative export price.

To be more precise, foreign income has a positive effect on exports, suggesting that a 10% rise in the incomes of OECD countries may increase the Turkish exports by 8.4%. Real exchange rate volatility affects the level of Turkish exports positively, implying that a complete elimination of the volatility of 2.58 would have decreased total exports by 4 percent. Third, as far as the short-run dynamics is concerned, about 74 percent of the variation in the Turkish exports is corrected within the next quarter.

In terms of policy implications, our study establishes that policy makers can not improve the contry's balance of trade both in the long-run and short-run by following price-oriented policies. Nor are policy makers required to follow an exchange rate stabilizing policy. Further research is also needed on disaggregated trade data to make clear policy recommendations.

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